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UNEMPLOYMENT INSURANCE AND THE LENGTH OF UNEMPLOYMENT. (U)

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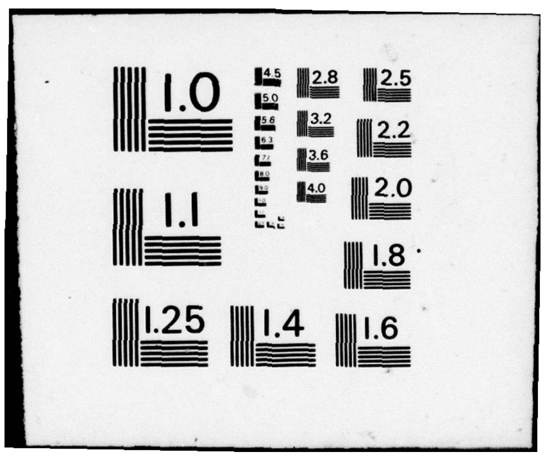
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UNEMPLOYMENT INSURANCE AND THE LENGTH OF UNEMPLOYMENT

Kathleen P. Classen

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UNEMPLOYMENT INSURANCE AND THE LENGTH OF UNEMPLOYMENT

INTRODUCTION

One explanation for persistently high rates of unemployment over the last two years is that unemployment insurance (UI) has reduced the incentives to return to work. There is no doubt that recent changes in UI laws have increased the potential work disincentive effects of UI. Coverage has been extended to virtually all wage and salary workers. Since many of the newly covered workers come from highly seasonal industries, they are frequently given the opportunity to prolong a spell of unemployment. Moreover, all of the new coverage is financed from federal revenues rather than through employer taxes. Thus, there is no reason for employers to protest when legally ineligible claimants file for benefits. These employers are, in fact, receiving a wage subsidy through UI payments. This subsidy increases employment in the newly covered seasonal industries.

The duration of benefits has also been extended from an average of 26 weeks to up to 65 weeks. Because of these changes in the UI law and because of the rapid changes in the demographic characteristics of the labor force, the noninflationary rate of unemployment (the "natural" rate) may be considerably higher than the target rates that are currently being suggested in both political and legislative arenas.

This paper presents evidence that UI does prolong the duration of unemployment for covered workers. Although this is only one of several channels through which UI could affect the unemployment rate, this evidence lends considerable weight to the argument that the effects of UI should be considered when the unemployment rate is used to judge the health of the economy or as an input in determining the appropriate stance of fiscal and monetary policy.[1]

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[1] UI may also affect the incidence and the industrial distribution of unemployment because of tax incentives (Feldstein [6]), and it may affect labor force participation and duration for noncovered workers (Mortensen [12]).

THE DATA

The data sets for this study come from two states: Arizona and Pennsylvania. They consist of quarterly earnings histories which are collected from employers to determine UI payroll taxes and to determine benefit entitlement when a claim is filed. These quarterly tax records were merged with UI claim histories to form a longitudinal work and claim history for a random sample of claimants in both states.

The results reported in this paper are based on claims filed in the year before and the year after an increase in benefit entitlement that occurred in Pennsylvania on Jan 1, 1968 and in Arizona on July 7, 1968. This two-year period was chosen to avoid seasonal bias and to provide variation in benefit entitlement that is independent of previous earnings.

The data that I used for this study allow me to avoid many of the problems that have been encountered in other studies of the effects of UI on the duration of unemployment.[2] A major advantage is that the data contain information on the benefit entitlement of claimants who returned to work immediately after filing a claim. Many of the estimates of the duration effect of UI that have been reported recently are based on survey data that records either the weekly benefit amount received or whether UI income was greater than zero. Since almost all states require a waiting period of one week, receipt of UI usually requires two weeks of unemployment. Those individuals who return to work before the end of the second week will not usually receive compensation. Any correlation between zero compensation and short durations for these individuals is not an indication of a work disincentive caused by UI. The link is actually reversed: short durations result in low (zero) compensation. Thus, studies that use benefit receipt, rather than benefit entitlement, will measure the duration effects of UI with an upward bias.

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[2] Estimates of the effect of UI on duration that are based on individual (as opposed to aggregate) data can be found in Lininger [9], Schmidt [13], Hills [7], Ehrenberg and Oaxaca [5], Holen [8] and Warner [15].

The data used in this study are also extremely accurate. Freedom from measurement error does not depend on the ability or willingness of a survey respondent to recall such key variables as earnings or the weekly benefit amount. Some of the information in surveys, particularly with respect to UI, can be extremely inaccurate.[3]

Another attractive feature of this data is that I was able to use every observation in the sample for the duration regressions. Survey data always requires that many observations be deleted because of measurement problems. The duration of unemployment, for example, cannot be measured when the respondent is unemployed on the survey date. Since the length of unemployment is positively related to the probability of being unemployed on the survey date, the observations are selected on the basis of the dependent variable, and hence, susceptible to bias. Since there is no need to eliminate observations for any reason in the sample results reported here, the probability that various selection criteria have introduced an unknown bias is reduced.

Claim histories do have a major drawback. They record only the compensated weeks of unemployment, not the total weeks. The measured duration of unemployment for claimants who exhaust their benefits is truncated at the maximum number of weeks of eligibility. There is, however, a maximum likelihood technique (TOBIT) that is designed for this type of truncation problem. Both OLS and TOBIT results are reported. They yield very similar parameter estimates, primarily because exhaustees account for only 8% of the observations in Pennsylvania and 14% of the observations in Arizona.

EMPIRICAL RESULTS

The hypothesis that an increase in UI leads to an increase in the duration of unemployment for covered workers can be derived either from a simple labor-leisure model, a simple search model (e.g. McCall [10] or Mortensen [11]), or a model that combines both leisure and search (e.g., Mortensen [12] or Burdett [1]).[4]

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[3] See Classen [4] for a discussion of the accuracy of UI responses in survey data.

[4] Mortensen [12] has pointed out that an increase in UI may decrease the expected duration for job seekers who are not covered <u>if</u> an increase in UI increases the expected return to being employed.

The appropriate form of the functional relationship between UI parameters and duration depends on the form of the utility function, and, in a search-theoretic framework, on the characteristics of the wage offer distribution and on what the searcher knows about the distribution. Since I cannot measure any of the above, I have investigated the robustness of the relationship between both the length and level of benefit payments and the duration of unemployment with respect to several specifications of the following general form:

$$D = f(\text{WBA}, \text{MAX WEEKS}, \bar{C}_i)$$

where

	D	= weeks of compensated unemployment
	WBA	= weekly benefit amount
	MAX WEEKS	= potential duration of benefits
and	$\bar{C}_i$	is a vector of personal and labor market characteristics

The weekly benefit amount (WBA) is determined by previous earnings and can be collected for up to the potential duration of benefits (MAX WEEKS) in the 52-week period following the date that the claim is filed (the benefit year). If more than one spell of unemployment occurs during the benefit year, the weekly benefit amount remains unchanged. Table I lists the UI legal parameters that determined WBA and MAX WEEKS for the two-year period covered in Pennsylvania (1967 and 1968) and the two-year period covered in Arizona (July 7, 1967 to July 7, 1969). No changes in the weekly benefit amount occurred in mid-claim due to an increase in the allowable maximum WBA.

The other independent variables that were included in the regression equations represent factors such as age and sex that are likely to be correlated with WBA and have an independent effect on the duration of unemployment. The two most important of these variables are earnings in the highest quarter of the base year (PREV. EARNINGS) and *the ratio of base year earnings to high quarter earnings* (EARNINGS STAB). The value of EARNINGS STAB must lie between 1.0 (earnings only in one quarter) and 4.0 (constant earnings over the base year). Taken together these two earnings variables measure

TABLE I

UI LEGAL PARAMETERS

	PENNSYLVANIA	ARIZONA
WBA	1/26 HIGH QUARTER EARNINGS Up to $45 in Year 1 Up to $60 in Year 2	1/25 HIGH QUARTER EARNINGS Up to $43 in Year 1 Up to $50 in Year 2
	(WBA in all cases is based on year in which claim was filed)	
MAX WEEKS	1/2 BASE PERIOD EARNINGS WBA Up to 30 weeks	1/3 BASE PERIOD EARNINGS WBA Up to 26 weeks

BASE PERIOD EARNINGS are the earnings in the first four of the last five calendar quarters preceding the date that the claim is filed and HIGH QUARTER EARNINGS are the earnings in the highest quarter of that base year.

the level and stability of previous earnings. They reflect the cost of being unemployed in terms of foregone earnings. EARNINGS STAB also measures the extent to which the duration of unemployment per year is correlated over time for individuals.

The industry dummies, which are based on the 2-digit SIC of the principal industry in the base year, are included to measure demand conditions relative to the omitted class (whole-retail trade). SICX means that the code was blank. The vast majority of these cases, which occur only in Arizona, refer to claimants who come from the military or the federal-civilian government. The Pennsylvania sample did not include any claims filed under federal programs. There are two cyclical variables in the regression equations. YEAR measures the effect of differences in cyclical conditions between the two sample years. Since any cyclical effect could differ by skill level (PREV. EARNINGS), a YEAR-SKILL interaction term was also included.[5]

Table II lists the coefficients from three different specifications of the relationship between the duration of unemployment and the independent variables described above. The first column under each state lists the estimated coefficients from an OLS linear regression. The second column is coefficients from a regression in which all continuous variables were measured in natural logs. Several other functional forms were also estimated. The two reported OLS functional forms (LINEAR and LN-LN) span the range of estimated WBA elasticities. The mean WBA elasticities for these two forms are listed in Table III.

All OLS regressions produced positive and significant estimates of the WBA coefficient with the elasticities falling within the ranges indicated in Table III. These results provide very strong evidence that the duration of unemployment (measured by compensated weeks) is positively related to the level of weekly benefit payments.

The coefficient of MAX WEEKS was also positive and significant in all the OLS regressions since the number of compensated weeks is limited by the number of weeks of potential duration. The positive coefficient does not necessarily imply that an

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[5]The exclusion of these cyclical variables does not affect the results reported in this paper.

TABLE II

DURATION REGRESSIONS

DEPENDENT VARIABLE = WEEKS OF COMPENSATED UNEMPLOYMENT

INDEPENDENT VARIABLES (t-statics in parentheses)	PENNSYLVANIA (N = 4240) Coefficients LINEAR OLS		LN-LN OLS		TOBIT		ARIZONA (N = 10,887) Coefficients LINEAR OLS		LN-LN OLS		TOBIT	
WBA	.11	(5.68)	1.03	(8.41)	.12	(5.79)	.12	(9.16)	.84	(11.37)	.14	(8.98)
MAX WEEKS	.19	(2.99)	1.09	(4.61)	.07	(.94)	.23	(9.18)	.58	(8.70)	.12	(3.82)
SEX (F=1)	.32	(.88)	.01	(.28)	.44	(1.11)	3.17	(16.26)	.36	(12.66)	3.77	(16.29)
AGE	.10	(10.69)	.46	(8.82)	.12	(11.10)	.13	(20.00)	.72	(19.00)	.15	(19.67)
PREV. EARNINGS	-.002	(-5.94)	-.60	(-7.28)	-.002	(-5.76)	-.001	(-8.57)	-.47	(-10.70)	-.001	(-8.12)
EARNINGS STAB	-3.07	(-10.83)	-1.32	(-12.36)	-3.13	(-10.05)	-1.25	(-8.76)	-.65	(-11.92)	-1.30	(-7.77)
INDUSTRY DUMMIES												
AGRICULTURE	1.55	(1.05)	.19	(.94)	1.66	(1.02)	.22	(.17)	.04	(.19)	.72	(.43)
MINING	-2.19	(-1.99)	-.25	(-1.68)	-2.45	(-2.04)	-2.26	(-6.34)	-.41	(-7.91)	-2.70	(-5.23)
CONSTRUCTION	1.03	(1.92)	.26	(3.61)	.94	(1.61)	.20	(.74)	.14	(3.56)	.31	(.97)
NON-DURABLES	-2.68	(-6.45)	-.33	(-5.83)	-2.97	(-6.55)	-.17	(-.48)	.08	(1.62)	-.14	(-.33)
DURABLES	-.53	(-1.18)	-.09	(-1.55)	-.61	(-1.25)	-.02	(-.06)	.05	(1.28)	.02	(.07)
TRANSPORTATION	-.79	(-.92)	-.14	(-1.23)	-1.00	(-1.08)	-2.01	(-5.11)	-.36	(-6.28)	-2.23	(-4.86)
FINANCE	3.50	(3.29)	.31	(2.16)	4.14	(3.51)	-.20	(-.41)	-.01	(-.09)	-.27	(-.47)
SERVICES	2.60	(4.18)	.31	(3.61)	3.08	(4.49)	.46	(1.61)	.11	(2.63)	.54	(1.61)
PUBLIC ADMIN.	–	–	–	–	–	–	-4.00	(-1.70)	-.55	(-1.61)	-4.80	(-1.80)
SIC X	–	–	–	–	–	–	1.26	(4.31)	.25	(6.02)	1.50	(4.45)
YEAR (YEAR 2 = 1)	-.32	(-.58)	1.07	(2.15)	-.36	(-.60)	-1.37	(-4.18)	-1.17	(-3.74)	-1.70	(-4.48)
YEAR-SKILL	-.000	(-.70)	-.16	(2.34)	-.000	(-.77)	.000	(2.49)	.15	(3.42)	.000	(2.62)
CONSTANT	5.76		-1.97		8.94		-2.70		-2.10		-0.88	
	R^2 = .09 F = 27.8		R^2 = .10 F = 29.3		χ^2 = 2534		R^2 = .09 F = 58.6		R^2 = .10 F = 67.5		χ^2 = 10,037	

TABLE III

WBA ELASTICITIES

	Pennsylvania	Arizona
LINEAR	.60	.61
LN-LN	1.03	.84

increase in potential duration results in an increase in the total duration of unemployment. Even if MAX WEEKS had no effect on total duration, there would still be a positive relationship between MAX WEEKS and compensated duration (CD).

If the relationship between total duration (TD) and the vector of all independent variables is linear, then

$$TD_i = \beta'\overline{X}_i + \mu_i$$

$$CD_i = TD_i \quad \text{if} \quad TD_i < \text{MAX WEEKS}_i$$

$$CD_i = \text{MAX WEEKS if} \quad TD \geq \text{MAX WEEKS}_i$$

OLS estimates that are based only on compensated duration ($CD_i = TD_i$ for all i) will consistently underestimate TD for exhaustees. This implies a systematic relationship between the error term and the exogenous variables. The OLS estimates are not, therefore, unbiased estimates of β'. In order to obtain efficient and consistent estimates of β' , I have used the technique described by Tobin [14] that maximizes the log of the following likelihood function:

$$L(\beta,\sigma/CD,\overline{X},MW) = \prod_{CD_i = MW_i} P\left(\frac{MW_i - \beta'\overline{X}_i}{\sigma}\right) \cdot \prod_{CD_i < MW_i} \frac{1}{\sigma} Z\left(\frac{MW_i - \beta'\overline{X}_i}{\sigma}\right)$$

where $Z(\cdot)$ is the standard normal density function and $P(\cdot)$ is the corresponding cumulative distribution function. MW is MAX WEEKS.

The third column under each state in Table II lists the TOBIT estimates of a linear duration equation. The major difference between the OLS results and the TOBIT results is in the coefficient of MAX WEEKS (compare columns 1 and 3 under each state). The TOBIT estimate is considerably lower and not significant for Pennsylvania. This is not surprising since the TOBIT equations provide estimates of $\frac{dTD}{dMW}$ and TD is not constrained by MAX WEEKS (MW). The estimated WBA coefficient is slightly higher for both states.

Although the sample period was chosen in each state to provide some variation in WBA that is independent of previous earnings, the WBA variation that occurs within years also provides very strong evidence that the level of benefit payments influences the duration of unemployment.[6] Within each sample year, the relationship between previous earnings and benefits is a constant fraction (α) of previous earnings (PE) up to the maximum (K). If we assume a linear functional form:

$$D = b_0 + b_1 WBA + b_2 PE + \Sigma b_i X_i \tag{1}$$

then

$$\begin{cases} D = b_0 + b_1 \alpha PE + b_2 PE + \Sigma b_i X_i & \text{if } PE < \frac{K}{\alpha} \\ D = b_0 + b_1 K + b_2 PE + \Sigma b_i X_i & \text{if } PE > \frac{K}{\alpha} \end{cases} \tag{2}$$

or

$$D = \gamma_0 + \gamma_1 PE + \gamma_2 [PE - \frac{K}{\alpha}] Z + \Sigma b_i X_i \tag{3}$$

where

$$Z = 1 \text{ if } PE > \frac{K}{\alpha}; \quad 0 \text{ otherwise}$$

$$\gamma_1 = \alpha b_1 + b_2$$

$$\gamma_2 = -\alpha b_1$$

[6] This test was suggested to me by Welch [16] in comments that he made on an earlier paper of mine [3].

Equation (3) is a spline function which allows the slope of the relationship between previous earning and duration to change at $\frac{K}{\alpha}$. Given equation (1), the slope will change ($\gamma_2 \neq 0$) only if $\frac{dD}{dWBA}$ (which is equal to b_1) is not zero. Figure 1 shows what the spline should look like according to the linear results in Table II.

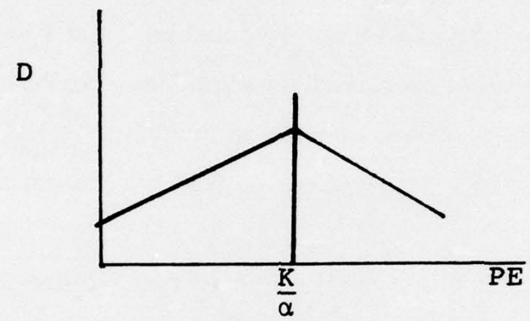

Figure 1

The problem with a spline such as figure 1 is that the relationship between duration and previous earnings, holding weekly benefit payments constant, may not be linear. The linear specification implies that the break at $\frac{K}{\alpha}$ is caused by the fact that benefits no longer increase with earnings. This break could, however, be the result of a non-linear relationship between PE and D that has been allowed to break only at $\frac{K}{\alpha}$.

In order to distinguish between these two explanations of the pattern shown in figure 1, I estimated a spline function with two break points. One break point is placed at the earnings level where benefits are equal to the maximum $(\frac{K}{\alpha})$ and one at an extraneous break point. If $\frac{dD}{dWBA}$ is positive, then the spline should break at $\frac{K}{\alpha}$, rather than at the extraneous break point.

The break point $(\frac{K}{\alpha})$ is different in the two years for both states. In order to make the graphs and computations somewhat simpler, the extraneous break point for year 1 was set equal to the $\frac{K}{\alpha}$ point in year 2 and vice versa. This leads to the following estimating equation:

$$D = \gamma_0 + \gamma_1 PE + \gamma_2 [PE - \frac{K_1}{\alpha}] Z_1 + \gamma_3 [PE - \frac{K_2}{\alpha}] Z_2 + \Sigma b_i X_i$$

where

$$Z_1 = 1 \text{ if } PE > \frac{K_1}{\alpha} \text{ ; 0 Otherwise}$$

$$Z_2 = 1 \text{ if } PE > \frac{K_2}{\alpha} \text{ ; 0 Otherwise}$$

$$\alpha = \frac{1}{26} \text{ for Penn; } \frac{1}{25} \text{ for Arizona}$$

The values of K_1 and K_2 are the WBA maximums shown in table I. The OLS estimates of γ_1 , γ_2 and γ_3 are presented in table IV. These equations included the independent variables listed in table II, with the exception of YEAR, YEAR-SKILL and WBA.

These regression estimates show a strong negative relationship between previous earnings and duration as soon as WBA stops increasing along with previous earnings. Since the relationship between PE and D is positive (or flat) when WBA and PE increase together, the effect of benefits on the duration of search must be positive. Figure 2 illustrates this relationship between duration and previous earnings. In year 1, WBA increases up to $\frac{K_1}{\alpha}$ and in year 2, WBA increases up to $\frac{K_2}{\alpha}$.

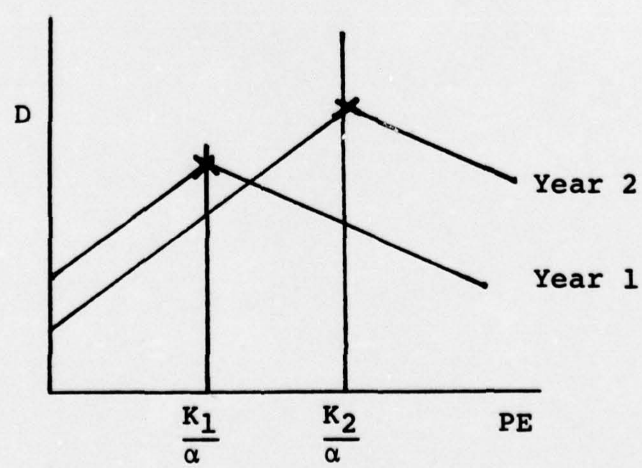

Figure 2

TABLE IV

TWO-BREAK SPLINE REGRESSIONS

Dependent Variable = Weeks of Compensated Unemployment
(t-statistics in parentheses)

	PENNSYLVANIA		ARIZONA	
	Year 1	Year 2	Year 1	Year2
Prev. Earnings (γ_1)	.002 (2.10)	-.000 (-.14)	.003 (4.20)	.004 (3.92)
Break 1 (γ_2)	$\boxed{-.006}$ $\boxed{(-2.17)}$	.005 (1.75)	$\boxed{-.008}$ $\boxed{(-2.88)}$	.002 (.64)
Break 2 (γ_3)	.001 (.57)	$\boxed{-.006}$ $\boxed{(-3.03)}$	.003 (1.50)	$\boxed{-.006}$ $\boxed{(-2.56)}$

CONCLUSION

provides

This paper ~~has provided very~~ strong evidence that UI benefits prolong the duration of unemployment and, therefore, that at least some unemployment is UI-induced. These results support the contention that the recent high rate of unemployment can be explained at least partially by the significant liberalizations that have occurred in the UI program over the last two years.

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